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A RECONSIDERATION OF THE EFFECTS OF UNIONISM ON RELATIVE WAGES AND EMPLOYMENT IN THE UNITED STATES, 1920-80

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ABSTRACT

H. Gregg Lewis' estimates of the relative wage effect of unionism between 1920 and 1958 are routinely cited though they have rarely been subject to scrutiny. This paper extends Lewis' data to 1980 and, in particular, we construct a series on union membership that links up with the data available in the 1970's from the Current Population Surveys. We proceed to reexamine the effects of trade unions both on relative wages and on relative manhours worked. Our estimates of the relative wage effect are similar to Lewis' though these are not measured with precision and a wide range of estimates are consistent with the results. With respect to the effect of unionism on relative manhours worked, we are not at all satisfied that the analysis of these data clearly points to the existence of a negative effect.

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A Reconsideration of the Effects of Unionism on Relative Wages and Employment in the United States, 1920-80

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I. Introduction

When an economist publishes important, original, and potentially controversial empirical results, it is usually the case that the data and the procedures generating these results are scrutinized by other economists. The chains of reasoning that led to these findings are examined, the data are inspected, and the estimates are checked to determine whether the results claimed do, indeed, follow and, if so, to ascertain the degree to which these results are sensitive to small changes in the underlying assumptions. It is, of course, this social character of the discipline which gives rise to any impartiality that economics can claim. As Popper (1966) expresses it, ". . . objectivity is closely bound up with the _social aspect of scientific method_, with the fact that science and scientific objectivity do not (and cannot) result from the attempts of an individual scientist to be "objective," but from the friendly-hostile cooperation of many scientists" (p. 217).

One study in economics that appears not to have been exposed to this collective scrutiny is H. Gregg Lewis' (1963, 1964) estimates of the relative wage and employment effects of U.S. trade unions over the years 1920-58. His results have justly received widespread attention, but they are cited in a routine manner and, for the most part, they have not been subject to the same critical evaluation accorded to other empirical studies of comparable standing in economics.1/ One of the very few assessments of Lewis' estimates was that provided by Melvin Reder (1965) in his insightful review of Lewis' book in the Journal of Political Economy for April 1965. In this paper, Reder not only supplied an excellent synopsis of Lewis' work, but also he clearly identified the critical points in Lewis' empirical procedures and he offered alternative explanantions for Lewis' findings. Indeed, some of the most original research in measuring the effects of trade unions on labor market variables are direct descendents of Reder's thoughtful review article.2/

The purpose of this paper is to resume the process of evaluating Lewis' work. We have extended Lewis' data and his estimates of the relative wage effects of unions from 1959 to 1980. We determine whether the inferences from Lewis' regressions are sensitive to alternative specifications of the equations including making some allowance for serial correlation in the estimated residuals. We also take up the question of the effects of unions on total manhours worked. Our procedure is first to specify what we call "descriptive equations" whose purpose is essentially to describe the underlying empirical regularities and where little recourse is made to structural labor supply and labor demand equations. We then specify a structural model, estimate the parameters of that model, and draw inferences from these estimates about the effects of unionism on relative wages and on relative manhours worked. We turn first to a brief description of the data and the variables that we shall be using in our empirical analysis.3/

II. Description of the Data

Lewis' purpose in Chapter VI of his book (1963) was to derive estimates of the average union-nonunion wage differential for the U.S. economy over the years from 1920 to 1958. Reduced to its essentials, his procedure was to divide the economy into two sectors and, after controlling for the effects of other variables, to compare the movements in the logarithms of relative wages in the two sectors with movements in the sectoral difference in the fraction of workers unionized. The two sectors were denoted by a and b: sector a consisted of the mining, construction, manufacturing, transportation, communications, and public utilities industries; sector b consisted of all other industries except for government work relief and the U.S. military. During the 1920-58 period examined by Lewis, more than 80 percent of the economy's union workers were employed in sector a.

The task of extending beyond 1958 the data on hourly compensation, manhours worked, and output in the two sectors is straightforward (if tedious) and, although there exist some differences between Lewis' procedures and our procedures in forming these variables, these differences are very small in our judgment. This is not the case, however, with our series on the fraction of workers covered by collective bargaining contracts in the two sectors and our task in constructing these series for the 1960's and 1970's was more formidable.[4/] (Our procedures in forming these variables are given in detail in the Appendix.) For these variables, our procedure was to start with the union coverage and union membership data collected by the Current Population Surveys (CPS) in the 1970's and then to link these data with those published biennually by the Bureau of Labor Statistics (BLS) in the 1960's and late 1950's. The resulting series are likely to be quite accurate for the 1970's, but will surely be substantially less so for the preceding years. We have tried to compile a series on the numbers of workers represented by trade unions rather than a series on union membership. In this respect, our concept is different from Lewis' although the difference between the sectors in the fraction of workers represented by unions (the variable to be used in most of the empirical analysis below) differs trivially from the sectoral difference in the fraction of workers who are union members, at least for the years from 1977 to 1980.[5]

Our estimates of the thousands of workers represented by trade unions from 1958 to 1980 are listed in Table 1: those for sector a are given in column (i), for sector b in column (ii), and for the entire economy in column (iii); the fraction of full-time equivalent employees represented by trade unions in sectors a and b is given in columns (iv) and (v) respectively of Table 1 and the arithmetic difference in the unionization proportions in the two sectors is given in column (vi). The well-known decline in the extent of unionism in the "old" industries is revealed by the unmistakable negative trend in the series in column (iv). This has been offset only slightly by the growth in the extent of unionism in sector b (primarily an expansion of unionism in government). As column (vi) of Table 1 makes clear, and is evident from Figure 1, the unionism difference between the two sectors narrowed considerably between 1958 and 1980: whereas the fraction of sector a's

Table 1

Extent of Union Representation, 1958-80

	Thousands of Workers Represented by Trade Unions			Fraction of Full-Time Equivalent Employees Represented by Trade Unions			
	Sector a	Sector b	Total	Sector a	Sector b	Difference	Total
	Column (i)	Column (ii)	Column (iii)	Column (iv)	Column (v)	Column (vi)	Column (vii)
1958	13,701.3	3,217.8	16,919.1	0.606	0.113	0.494	0.331
1959	13,725.5	3,185.3	16,910.8	0.585	0.109	0.477	0.320
1960	13,749.6	3,152.8	16,902.4	0.584	0.104	0.480	0.314
1961	13,478.3	3,160.6	16,638.9	0.587	0.103	0.484	0.311
1962	13,206.9	3,168.3	16,375.2	0.560	0.101	0.459	0.298
1963	13,234.5	3,365.6	16,600.1	0.557	0.105	0.452	0.298
1964	13,262.0	3,562.8	16,824.8	0.550	0.108	0.441	0.295
1965	13,618.4	3,757.0	17,375.4	0.538	0.110	0.428	0.293
1966	13,974.8	3,951.1	17,925.9	0.522	0.111	0.411	0.287
1967	13,901.0	4,791.6	18,692.6	0.514	0.130	0.385	0.292
1968	13,827.3	5,632.0	19,459.3	0.503	0.147	0.355	0.296
1969	13,497.8	5,689.3	19,187.1	0.479	0.144	0.335	0.284
1970	13,168.3	5,746.6	18,914.9	0.483	0.143	0.341	0.280
1971	13,081.6	6,027.6	19,109.2	0.494	0.147	0.348	0.283
1972	12,994.9	6,308.5	19,303.4	0.478	0.148	0.330	0.277
1973	13,315.7	6,669.0	19,984.7	0.466	0.150	0.316	0.274
1974	13,265.7	6,888.7	20,154.4	0.465	0.151	0.314	0.272
1975	11,617.8	7,018.3	18,636.1	0.444	0.153	0.292	0.258
1976	11,963.3	7,372.5	19,335.8	0.442	0.156	0.287	0.260
1977	12,387.4	6,905.7	19,293.1	0.440	0.141	0.299	0.250
1978	12,582.6	7,019.6	19,602.2	0.425	0.137	0.288	0.242
1979	13,382.8	7,642.2	21,025.0	0.436	0.144	0.293	0.251
1980	12,386.9	7,535.4	19,922.3	0.417	0.140	0.277	0.238

employees represented by unions was almost 50 percent greater in 1958 than the fraction of sector b's employees, by 1980 this difference had almost halved to 28 percent. A difference between the two sectors of this order of magnitude had not been recorded since the late 1930's. (See Lewis (1963), Table 53.) However, in the late 1930's, well over 80 percent of all union workers were employed in sector a; by comparison, the corresponding figure in 1980 was a little over 62 percent.[6/]

The mean values and standard deviation of the variables used in the analysis are given in Table 2. The sample data we use for the period 1920-58 are identical to those used by Lewis except that for several variables we altered his values for 1958 (the last observation in his study) to take advantage of revisions in the figures by various government agencies.[7/] These alterations were small and reestimating Lewis' equations with these revised values for 1958 had trivial effects on his estimated parameters. As is evident in Table 2, the size of sector a relative to sector b was smaller during the 1959-80 period than in the 1920-58 period whether size is measured in terms of labor input (total manhours) or in terms of output (national income). As we have already noted, the fraction of employees represented by unions in sector a was declining from 1958 to 1980 while that in sector b was growing slightly. Over the same period, the hourly compensation in sector a relative to that in sector b traces a U-shape declining from 1.43 in 1958 to 1.29 in 1970 and then rising to 1.37 in 1980, the same value as that in 1955. (See Figure 1) Although we shall sometimes refer to the hourly compensation series as a relative wage series, it

Table 2

Mean Values (and Standard Deviations in Parentheses) of Variables 1920-80, 1920-58, and 1959-80

	1920-80	1920-58	1959-80
W, average hourly compensation in sector a divided by average hourly compensation in sector b	1.325 (0.075)	1.312 (0.086)	1.348 (0.039)
E, total manhours worked in sector a divided by total manhours worked in sector b	0.903 (0.163)	0.974 (0.154)	0.775 (0.080)
P_a, the proportion of full-time equivalent employees unionized in sector a	0.421 (0.146)	0.377 (0.163)	0.499 (0.055)
P_b, the proportion of full-time equivalent employees unionized in sector b	0.085 (0.044)	0.060 (0.030)	0.131 (0.020)
$P = P_a - P_b$	0.335 (0.117)	0.317 (0.133)	0.368 (0.073)
Q, national income originating in sector a divided by national income originating in sector b	1.042 (0.156)	1.069 (0.177)	0.994 (0.092)
X, the ratio of the actual to the "expected" price level	1.046 (0.066)	1.029 (0.072)	1.076 (0.039)
Z, the unemployment rate of the entire labor force	0.071 (0.057)	0.082 (0.069)	0.052 (0.012)

NOTES: To facilitate comparison with Lewis' work, we adopt notation that is very similar to his.

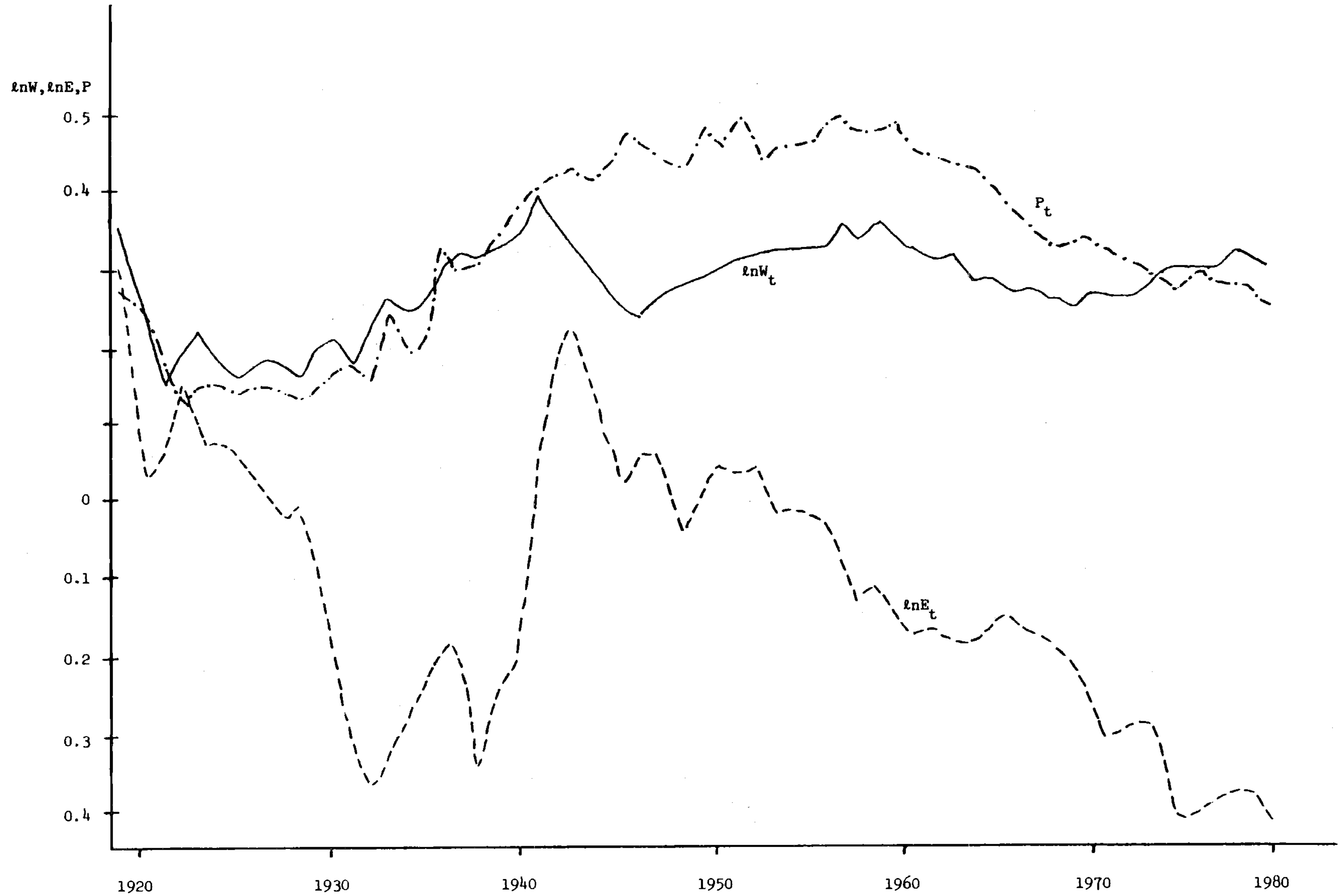

Figure 1: The Logarithm of Relative (ℓnW), the Logarithm of Relative Manhours Worked (ℓnE_t), and the Sectoral Difference in Unionization (P_t)

should be emphasized that these data include not merely wages and salaries, but also commissions, tips, bonuses, the value of payments in kind, estimated employer contributions to pension and health plans, and compensation for injuries. However, as was the case with Lewis' study from 1920 to 1958, these compensation figures have not been adjusted for changes in relative labor force composition except for shifts of employment among large industry aggregates. Therefore, some of the movements in these relative compensation figures will be attributable to changes in the relative "quality" of the employees and, by not accounting for these changes in "quality," our procedures are likely to overestimate the true impact of unionism on relative wages. For, insofar as unionized employers retain considerable discretion over employment decisions, they are induced by a higher union wage to upgrade the "quality" of their employees so that a part of what we measure as the relative wage effect of unionism is likely to be, in fact, a wage difference that corresponds to differences in the "quality" of the work forces of unionized employers compared with those of nonunionized employers. This was a limitation of Lewis' study and it applies also to this one.

The other two variables listed in Table 2, X_t and Z_t, describe two features of the economy at the aggregate level. Both of these variables were used by Lewis in his analysis. X_t is the ratio of the price level to the value implied by the recent trend in prices. That is, if p_t is the price level in year t and p_t^* is a weighted average of current and past values of p_t where the weights decline exponentially from the current period to earlier periods, then $X_t = p_t/p_t^*$.[8/] This variable takes on its highest values not so much in periods of inflation, but in periods of accelerating inflation (that is, when prices in year t are much higher than their recent values) and these occur in 1920, 1942-44, 1947-48, and 1974-80; X_t takes on its lowest values when prices in year t fall below the levels recorded in years immediately preceding year t and the extreme values are recorded in the early 1930's. Z_t measures the unemployment rate in the labor force.[9/] Once again, the post-1958 period displays marked constrast to the 1920-58 period: whereas the logarithm of X_t (which is the form in which X_t is used in the analysis below) and Z_t were negatively correlated in the years from 1920 to 1958, they were positively correlated during the period 1959-80.[10/]

III. Descriptive Equations

A. Wage Rates

Let the average hourly compensation received by workers covered by collective bargaining contracts (hereafter called "the union wage") within sector i during period t be given by W_{it}^u and let the average hourly compensation received by workers not covered by collective bargaining contracts (hereafter called "the nonunion wage") within sector i during period t be given by W_{it}^n. Then, during period t, the logarithm of the average wage observed in sector i, $\ell n\ W_{it}$, may be expressed as follows:[11/]

$$(1) \qquad \ell n\ W_{it} = \ell n\ W_{it}^{n} + [\ell n(1 + r_{it})]P_{it} = \ell n\ W_{it}^{n} + B_{it}P_{it} \quad ,$$

where i equals a or b, where P_{it} measures the fraction of workers within sector i covered by collective bargaining contracts (hereafter called "the fraction unionized") during period t and where $r_{it} = (W_{it}^{u} - W_{it}^{n})/W_{it}^{n}$ is the proportional difference within sector i during period t between the union and nonunion wage rates (hereafter called "the relative wage effect of unionism").[12] Subtract equation (1) for sector b from that for sector a:

$$(2) \qquad \ell n\ W_{t} = \ell n\ W_{t}^{n} + B_{at}P_{at} - B_{bt}P_{bt} \quad ,$$

where $\ell n\ W_{t} = \ell n(W_{at}/W_{bt})$ and $\ell n\ W_{t}^{n} = \ell n(W_{at}^{n}/W_{bt}^{n})$. A natural starting point in an empirical analysis is to investigate the consequences of assuming that the relative wage effects of unionism within each sector are constants over time. In this case, writing $\ell n\ W_{t}^{n} = \theta Y_{t} + u_{t}$ where Y_{t} is a vector of exogenous variables, θ a corresponding vector of unknown parameters, and u_{t} a stochastic error term, equation (2) may be written:

$$(3) \qquad \ell n\ W_{t} = \theta Y_{t} + B_{a}P_{at} - B_{b}P_{bt} + u_{t} \quad .$$

This equation was fitted to the annual data described in Section II over the period 1920-80 both by ordinary least-squares and by a generalized least-squares procedure that adjusts the variables for first-order serial correlation in the estimated residuals. Y_{t} was represented by a number of different variables including the logarithm of relative output in the two sectors $(\ln Q_t)$, the logarithm of relative wages lagged one year $(\ln W_{t-1})$, the logarithm of the ratio of the price level in period t to its level in years in and immediately before year t $(\ln X_t)$, the unemployment rate (Z_t), the logarithm of relative manhours worked lagged one year $(\ln E_{t-1})$, and a linear time trend (T_t). A representative sample of estimates are given in Table 3 where, from line to line, the point estimates of B_a and B_b change substantially (measured in terms of their implications for r_a and r_b).[13] However, these coefficients (especially B_b) are not estimated very precisely and a wide range of values for B_a and B_b are consistent with the data. Indeed, for most of the specifications we fitted based on equation (3), we could not reject the null hypothesis of no difference between the estimates of B_a and B_b. This provides some support for Lewis' procedure of focusing upon the arithmetic difference between P_{at} and P_{bt}, as we now proceed to do.

Let $\bar{B}_t$ be the following transformation of the economy-wide average of the proportionate union-nonunion wage differential: $\bar{B}_t = q_t B_{at} + (1 - q_t)B_{bt}$ where q_t measures the fraction of all unionized workers during period 5 who are employed in sector a. Then equation (2) may be written as follows:

$$(4)\ \ln W_t = \ln W_t^n + \bar{B}_t(P_{at} - P_{bt}) + (B_{at} - \bar{B}_t)P_{at} + (\bar{B}_t - B_{bt})P_{bt} \quad .$$

Now in this equation, the final term, $(\bar{B}_t - B_{bt})P_{bt}$, is likely to constitute a relatively small component of the movements in the entire right-hand side because, over the period 1920-80, P_{bt} averaged 0.085

Table 3

Parameter Estimates (with Estimated Standard Errors in Parentheses) of Equation (3)

Equation Number	Estimation Method	$\hat{B}_a$	$\hat{B}_b$	Estimated Coefficients on: Constant	$\ln Q_t$	$\ln E_{t-1}$	$T_t/100$	$\ln W_{t-1}$	$\ln X_t$	Z_t	$\hat{\rho}$	d/h	see	R^2
(3a)	OLS	0.360 (0.049)	0.278 (0.163)	0.152 (0.015)								0.47	0.037	0.599
(3b)	COOR	0.424 (0.081)	0.288 (0.308)	0.122 (0.030)							0.73 (0.09)	1.54	0.023	0.388
(3c)	OLS	0.266 (0.079)	0.526 (0.569)	0.156 (0.022)	0.132 (0.056)	-0.119 (0.045)	0.034 (0.144)	0.108 (0.094)				6.47	0.033	0.696
(3d)	COOR	0.154 (0.057)	0.524 (0.392)	0.082 (0.018)	0.104 (0.038)	-0.119 (0.031)	0.066 (0.099)	0.500 (0.081)			0.06 (0.13)	0.99	0.025	0.852
(3e)	OLS	0.135 (0.069)	-0.260 (0.229)	0.129 (0.015)	0.232 (0.060)				0.360 (0.088)	0.698 (0.105)		0.68	0.027	0.797
(3f)	COOR	0.250 (0.082)	0.001 (0.324)	0.105 (0.027)	0.152 (0.061)				0.413 (0.121)	0.659 (0.157)	0.69 (0.09)	1.91	0.020	0.603

NOTES: The mean of the dependent variable over the period 1920-80 is 0.280 and its standard deviation is 0.057. The standard error of estimate of the regression equation is given by see. OLS stands for ordinary least-squares and COOR for Cochrane and Orcutt's iterative procedure to adjust for first-order serial correlation in the fitted residuals. The final value of the first-order serial correlation parameter as estimated by COOR is given by $\hat{\rho}$. The Durbin-Watson statistic (d) is reported in equations (3a), (3b), (3e), and (3f) and Durbin's h statistic is reported for equations (3c) and (3d). When applying COOR, the statistics reported in the final three columns are based on the $\hat{\rho}$-adjusted variables and, therefore, they are not strictly comparable with those corresponding to the OLS estimates.

and was always less than 0.16. Moreover, the third term on the right-hand side of equation (4), $(B_{at} - \bar{B}_t)P_{at}$, is equal to $(B_{at} - B_{bt})(1 - q_t)P_{at}$ and this will also constitute a relatively small part of the movements of the entire right-hand side insofar as B_{at} differs little from B_{bt} and in view of the fact that $1 - q_t$ was less than 0.2 until 1963. Consequently, if the sum of the last two terms on the right-hand side of equation (4) is denoted by v_t^*, we may write

$$(5) \qquad \ln W_t = \theta Y_t + \bar{B}_t P_t + v_t^* + u_t$$

where, as before, $\ln W_t^n = \theta Y_t + u_t$ and where $P_t = P_{at} - P_{bt}$. Of course, equation (5) follows from equation (4) exactly (except for v_t^*) if $B_{at} = B_{bt} = \bar{B}_t$.

Once again, first consider the simplest assumption according to which $\bar{B}_t$ is a constant, $\bar{B}$:

$$(6) \qquad \ln W_t = \theta Y_t + \bar{B} P_t + v_t \quad ,$$

where $v_t = v_t^* + u_t$. Table 4 presents the consequences of estimating equation (6) to annual time-series observations from 1920 to 1980 where, according to the specification, Y_t is represented by $\ln Q_t$, $\ln W_{t-1}$, $\ln E_{t-1}$, $\ln X_t$, and Z_t. For each equation specification, we present a pair of estimates, one corresponding to conventional least-squares estimates and the other to estimates that allow for first-order serial correlation in the fitted residuals. As is evident from Table 4, the ordinary least-squares fitted residuals clearly exhibit positive serial

Table 4

Parameter Estimates (with Estimated Standard Errors in Parentheses) of Equation (6)

Equation Number	Estimation Method	Estimated Coefficients on: Constant	P_t	ℓnQ_t	ℓnX_t	Z_t	ℓnE_{t-1}	ℓnW_{t-1}	$\hat{\rho}$	d/h	see	R^2
(6a)	OLS	0.153 (0.014)	0.378 (0.040)							0.46	0.037	0.596
(6b)	COOR	0.127 (0.028)	0.446 (0.075)						0.75 (0.09)	1.55	0.023	0.377
(6c)	OLS	0.128 (0.015)	0.259 (0.046)	0.125 (0.040)	0.459 (0.080)	0.585 (0.097)				0.56	0.028	0.777
(6d)	COOR	0.113 (0.026)	0.294 (0.072)	0.119 (0.050)	0.467 (0.110)	0.622 (0.154)			0.73 (0.09)	1.94	0.020	0.580
(6e)	OLS	0.123 (0.018)	0.170 (0.053)	0.175 (0.044)	0.411 (0.078)	0.502 (0.095)	-0.055 (0.027)	0.130 (0.072)		4.98	0.026	0.813
(6f)	COOR	0.247 (0.073)	0.244 (0.089)	0.196 (0.051)	0.328 (0.115)	0.912 (0.166)	0.175 (0.039)	0.175 (0.039)	0.96 (0.04)	0.56	0.018	0.656

NOTES: See the notes beneath Table 3. The Durbin-Watson statistic (d) is presented for equations (6a), (6b), (6c), and (6d) and Durbin's h statistic is presented for equations (6e) and (6f).

correlation although, in almost every case, the point estimates of the parameters are not sensitive to the modelling of a simple first-order autoregressive error process. For every specification in Table 4 and in many others whose estimates are not reported, the null hypothesis of a zero relative wage effect of unionism can be rejected at a high level of significance. The estimated coefficient on P_t falls as more variables are added to the regression equation so that, in equations (6e) and (6f), the implied point estimates of r are 18.5 percent and 27.6 percent respectively. These results for the period 1920-28 are similar to Lewis' results for 1920-58 when he fits equations that assume $\bar{B}_t$ is constant.[14]

Consider now the consequences of permitting the relative wage effect of unionism to vary over the years from 1920 to 1980. Lewis conjectured that unionism tended to reduce the responsiveness of money wages of union labor to transitory changes in the general price level and in employment. If this is the case, $\bar{B}_t$ would fall when such transitory elements are positive and $\bar{B}_t$ would rise when transitory elements are negative; that is, $\bar{B}_t$ would vary with the business cycle, being largest in the contraction phase and being smallest in the expansionary phase of the cycle. To implement this hypothesis, return to equation (5) and express $\bar{B}_t$ as a function, first, of the logarithm of the ratio of the price level in period t to its level in and immediately prior to year t ($\ell n\ X_t$) and, second, of the unemployment rate (Z_t): $\bar{B}_t = b_0 + b_1\ \ell n\ X_t + b_2 Z_t$ where $b_1 < 0$ and $b_2 > 0$. Substituting this expression for $\bar{B}_t$ into equation (5), we have:

$$\ell n\ W_t = \theta Y_t + b_0 P_t + b_1(P_t\ \ell n\ X_t) + b_2(P_t Z_t) + v_t \quad , \tag{7}$$

where $v_t = v_t^* + u_t$. Table 5 presents the consequences of fitting this equation to the annual data from 1920 to 1980 where again Y_t is given by $\ell n\ Q_t$, $\ell n\ X_t$, Z_t, and $\ell n\ E_{t-1}$. The null hypothesis that the value of $\bar{B}_t$ is constant over the 1920-80 period is easily rejected.[15/] For each equation in Table 5, b_1 is negative as conjectured and significantly less than zero by conventional criteria while b_2 is negative instead of the hypothesized positive sign. (If a two-tailed test were applied to b_2, then by conventional criteria it would not be judged significantly different from zero.) Moreover, simply evaluating the point estimates of b_1 and b_2, $\bar{B}_t$ is more responsive to movements in the inflation variable than to movements in unemployment: for instance, using equation (7d)'s estimates, estimated at the sample mean values of $\ell n\ X_t$ and Z_t, a one standard deviation increase in $\ell n\ X_t$ (with the unemployment rate constant) reduces the implied value of $\bar{B}_t$ from 23.1 percent to 9.1 percent whereas a one standard deviation increase in Z_t (holding the inflation variable constant) reduces the implied value of $\bar{B}_t$ from 23.1 to 16.6 percent. These results replicate Lewis' findings that unionism tended to make the money wages of union labor less responsive to sudden movements in the general price level while the evidence associating unionism's relative wage impact to the unemployment rate is considerably weaker.

The four equations in Table 5 imply similar movements in $\bar{B}_t$ over

Table 5

Parameter Estimates (with Estimated Standard Errors in Parentheses) of Equation (7)

Equation Number	Estimation Method	Estimated Coefficients on: Constant	P_t	$P_t \ell nX_t$	$P_t Z_t$	ℓnQ_t	ℓnX_t	Z_t	ℓnE_{t-1}	$\hat{\rho}$	d	see	R^2
(7a)	OLS	0.105 (0.019)	0.399 (0.071)	-3.566 (0.789)	-1.140 (0.958)	0.149 (0.033)	1.434 (0.235)	0.899 (0.257)			0.85	0.023	0.859
(7b)	COOR	0.088 (0.029)	0.434 (0.103)	-2.568 (0.947)	-1.195 (1.018)	0.110 (0.045)	1.177 (0.288)	0.903 (0.281)		0.65 (0.10)	1.90	0.019	0.668
(7c)	OLS	0.114 (0.020)	0.362 (0.075)	-3.117 (0.834)	-0.832 (0.951)	0.176 (0.039)	1.261 (0.256)	0.782 (0.261)	-0.028 (0.025)		0.95	0.023	0.859
(7d)	COOR	0.099 (0.039)	0.405 (0.132)	-2.177 (0.977)	-1.136 (1.034)	0.115 (0.051)	1.128 (0.051)	0.992 (0.279)	0.086 (0.036)	0.82 (0.07)	1.89	0.018	0.633

NOTES: See notes beneath Table 3.

Table 6

Point Estimates (and Estimated Standard Errors in Parentheses) of $\bar{B}_t$

Period	Column (i) Equation (7c)	Column (ii) Equation (7d)	Column (iii)	Column (iv) Structural Estimates
1920-24	0.225 (0.043)	0.273 (0.082)	0.276 (0.070)	0.161 (0.065)
1925-29	0.347 (0.055)	0.375 (0.109)	0.362 (0.093)	0.255 (0.060)
1930-34	0.496 (0.109)	0.393 (0.120)	0.432 (0.119)	0.418 (0.090)
1935-39	0.223 (0.114)	0.210 (0.113)	0.256 (0.120)	0.188 (0.081)
1940-44	0.049 (0.070)	0.150 (0.090)	0.161 (0.088)	0.015 (0.078)
1945-49	0.007 (0.068)	0.136 (0.094)	0.138 (0.090)	-0.027 (0.080)
1950-54	0.148 (0.044)	0.235 (0.085)	0.231 (0.073)	0.090 (0.067)
1955-59	0.203 (0.041)	0.269 (0.085)	0.266 (0.071)	0.139 (0.064)
1960-64	0.210 (0.040)	0.275 (0.084)	0.274 (0.070)	0.153 (0.064)
1965-69	0.202 (0.040)	0.258 (0.087)	0.252 (0.073)	0.117 (0.065)
1970-74	0.079 (0.056)	0.160 (0.088)	0.165 (0.084)	0.013 (0.077)
1975-80	-0.064 (0.096)	0.071 (0.110)	0.114 (0.075)	-0.079 (0.090)
1920-80	0.170 (0.051)	0.231 (0.080)	0.239 (0.072)	0.117 (0.069)

the 1920-80 period. Columns (i) and (ii) of Table 6 present the point estimates and standard errors of $\bar{B}_t$ corresponding to equations (7c) and (7d) averaged over the values of $\ln X_t$ and Z_t in each subperiod and in the 1920-80 period as a whole. According to these estimates, the average relative wage effect of unionism, r, over the entire period was between 18.5 percent and 26.0 percent, values virtually identical to those implied by equations (6e) and (6f) in Table 4. The point estimates of r range widely over the six decades from a maximum in the first half of the 1930's to a minimum in the second half of the 1970's. These fluctuations tend to be slightly greater when measured by conventional least-squares than when estimated by generalized least-squares. However, the standard errors attached to the point estimates of $\bar{B}_t$ caution against any confident statements about the precise magnitude of the relative wage effect in any subperiod: for instance, according to equation (7d), a 95 percent confidence interval on $\bar{B}_t$ ranges from 15.3 percent to 88.3 percent in the early 1930's and from -14.9 percent to 29.1 percent in the late 1970's. That these union wage effects in particular subperiods are estimated very imprecisely was noted by Lewis, but this fact is typically overlooked by others reporting these estimates.[16]

The point estimates of $\bar{B}_t$ in Table 6 for the 1960's do not differ appreciably from those for the 1950's, but those in the 1970's are noticeably lower. In view of the relatively high values taken by the inflation variable, $\ln X_t$, in the 1970's, these lower values of $\bar{B}_t$ in this decade should not occasion much surprise. The disturbing aspect of this result is that it does not appear to conform to the measurement of the relative wage effects of unionism as estimated from various large data sets on individuals. For instance, Ashenfelter's (1978) analysis of individuals in the CPSs suggests an average value of $\bar{B}$ that rises from 11.6 percent in 1967 to 14.8 percent in 1973 and to 16.8 percent in 1975. Or Moore and Raisian's (1983) analysis of the Income Dynamics Panel for male heads of households suggests either a rising trend or no trend (depending upon the particular equation specification) in the union-nonunion wage differential from 1967 to 1977. George Johnson (1983) presents more evidence against the notion of a falling union relative wage effect in the 1970's. The changes estimated in these studies are not always measured very precisely, nor are those estimated in Table 6, so not too much should be made of this difference. Nevertheless, insofar as our point estimates of $\bar{B}_t$ in the 1970's contradict those inferred from the analysis of large data sets on individual workers, we are inclined to place more credence in the union wage effects estimated from the latter than in those estimated from highly aggregated data such as those in this paper.

One possible reconciliation that we considered was that the relationship between $\bar{B}_t$ on the one hand and the inflation variable and the unemployment rate on the other hand differed in the 1920-58 period that Lewis analyzed from that obtaining after 1958. After all, we have already noted in Section II that the covariance between $\ell n\ X_t$ and Z_t is sharply different in one period compared with the other. In fact, a conventional test of the null hypothesis that b_1 and b_2 are the same in the years 1959-80 as in the years 1920-58 can be rejected.[17] Consider, therefore, measuring $\bar{B}_t$ according to the following specification where D_t is a dummy variable taking the value of unity during the years from 1959 to 1980 and taking the value of zero otherwise: $\bar{B}_t = b_0 + b_1 \ln X_t + b_2 Z_t + b_3 D_t \ln X_t + b_4 D_t Z_t$. When this expression for $\bar{B}_t$ is substituted into equation (5) and when Y_t is represented by a linear combination of $\ln Q_t$, $\ln X_t$, Z_t, and $\ln E_{t-1}$, the resulting generalized least-squares estimates of $\bar{B}_t$ are those given in column (iii) of Table 6. Clearly, even allowing for a different structure for $\bar{B}_t$ in the post-1958 period, our point estimates of unionism's relative wage effect remain unchanged and, in particular, we continue to estimate a lower relative wage effect in the 1970's compared with the preceding decade.[18]

To conclude this section, we report briefly the consequences of addressing two other issues. First, we examined the effects of treating unionism and relative output as endogenous in equation (7). The unionism variable, P, almost certainly contains measurement error while the relative output produced in sectors a and b, $\ln Q$, is surely a function, in part, of the relative wages in the two sectors: for both P and $\ln Q$, there exists good reason for considering instrumental variable estimates of the relative wage equation.[19] The results were not at variance with those reported in Tables 5 and 6 because the standard errors of the resulting estimates were so large as to encompass a very wide range of different values including the point estimates of $\bar{B}_t$ already reported. This should serve to underline the uncertainty that surrounds the point estimates of $\bar{B}_t$ in particular subperiods.[20/]

Second, we took up Reder's (1965) "one substantive criticism" of Lewis' regression specifications. Reder writes, "It is unlikely that the effect of unemployment on relative wage rates was constant throughout the period. In the 1930's when unemployment was very high, New Deal legislation and political intervention worked substantially to raise wages in the union sector relative to the non-union, thereby giving a downward bias to the estimate of the effect of unemployment on the ratio of union to non-union wages." We implemented Reder's hypothesis by creating a dummy variable taking the value of unity from the years 1932 to 1941 inclusive. Then, first, we added this variable to the vector Y_t in estimating equation (7) and, second, in addition, we allowed b_2 to be a function of this dummy variable. Estimating these equations left our inferences from Tables 5 and 6 unaltered and the hypothesis that the effect of unemployment on $\bar{B}_t$ was the same in the 1930's as in other decades could not be rejected.[21/]

B. Manhours Worked

The empirical analysis reported above concerned the effect of unionism on relative wages.[22/] The results strongly suggest a positive union-nonunion wage differential although the magnitude of this differential is not estimated with any useful precision. Now, according to one popular characterization of the determination of wages and employment in unionized markets, any wage increases effected by trade unions will be associated with decreases in the utilization of the labor input. Therefore, it seems natural to enquire whether, over the period

1920-80, movements of total manhours worked in sector a relative to sector b are negatively associated with movements in the difference in unionism in the two sectors. For this purpose define $\bar{C}_t$ as the effect on the logarithm of manhours in sector a relative to sector b of an increase in the difference between sectors a and b in the fraction of workers unionized. In other words, what $\bar{B}_t$ is to wages, $\bar{C}_t$ is to manhours. Then, consider the following equations which are the manhours counterparts to equations (6) and (7) above:

$$(8) \qquad \ln E_t = \mu Y_t + \bar{C} P_t + v_{1t}$$

$$(9) \qquad \ln E_t = \mu Y_t + c_0 P_t + c_1 (P_t \ln X_t) + c_2 (P_t Z_t) + v_{1t}$$

where v_{1t} is a stochastic disturbance term and μ is a vector of parameters corresponding to the vector of variables Y_t. In equation (8), $\bar{C}_t$ is assumed to be a constant $\bar{C}$ whereas in equation (9) $\bar{C}_t$ is assumed to be a constant $\bar{C}$ whereas in equation (9) $\bar{C}_t$ varies over time according to the expression $\bar{C}_t = c_0 + c_1 \ln X_t + c_2 Z_t$.[23] The results from estimating equations (8) and (9) to the annual observations from 1920 to 1980 are presented in Table 7 with the implied values over time of $\bar{C}_t$ for equations (9c) and (9d) given in columns (i) and (ii) of Table 8.[24]

In the estimates based on equation (8), unionism exerts a negative effect on relative manhours worked in three out of four instances. The coefficient estimates on P_t vary considerably from equation to equation and in those equations that allow for first-order serial correlation in the residuals we cannot reject the null hypothesis that there is

Table 7

Parameter Estimates (with Estimated Standard Errors in Parentheses) of Equations (8) and (9)

Equation Number	Estimation Method	Estimated Coefficients on: Constant	P_t	$P_t \ell nX_t$	$P_t Z_t$	ℓnQ_t	ℓnX_t	Z_t	ℓnE_{t-1}	$\hat{\rho}$	d/h	see	R^2
(8a)	OLS	0.078 (0.023)	-0.433 (0.074)			0.648 (0.070)			0.650 (0.042)		1.57	0.044	0.939
(8b)	COOR	-0.158 (0.072)	-0.156 (0.156)			0.766 (0.059)			0.069 (0.067)	0.90 (0.06)	1.01	0.034	0.766
(8c)	OLS	0.186 (0.055)	-0.842 (0.163)			1.262 (0.142)	-0.483 (0.283)	-0.557 (0.345)			0.34	0.100	0.709
(8d)	COOR	-0.220 (0.077)	0.012 (0.125)			0.384 (0.074)	0.047 (0.152)	-1.374 (0.234)		0.95 (0.04)	1.74	0.038	0.866
(9a)	OLS	0.080 (0.023)	-0.519 (0.088)	0.608 (0.428)	0.794 (0.589)	0.666 (0.071)			0.673 (0.043)		1.04	0.044	0.945
(9b)	COOR	-0.405 (0.071)	0.500 (0.137)	0.454 (0.393)	-5.513 (0.652)	0.365 (0.052)			-0.039 (0.041)	0.95 (0.04)	0.94	0.021	0.919
(9c)	OLS	0.107 (0.036)	-0.664 (0.137)	3.904 (1.521)	1.904 (1.745)	0.693 (0.070)	-1.016 (0.464)	-0.344 (0.479)	-1.016 (0.464)		1.61	0.042	0.951
(9d)	COOR	-0.440 (0.088)	0.488 (0.179)	0.011 (1.132)	-4.884 (1.197)	0.297 (0.063)	0.142 (0.344)	-0.385 (0.301)	-0.058 (0.043)	0.96 (0.04)	0.58	0.020	0.927

NOTES: The mean of the dependent variable over the 1920-80 period is -0.113 with a standard deviation of 0.174. The Durbin-Watson statistic (d) is reported in equations (8c) and (8d) and Durbin's h statistic is reported for all other equations. Also see the notes beneath Table 3.

Table 8

Point Estimates (and Estimated Standard Errors in Parentheses) of $\bar{c}_t$

Period	Column (i) Equation (9c)	Column (ii) Equation (9d)	Column (iii) Structural Estimates
1920-24	-0.436 (0.079)	0.174 (0.110)	-0.113 (0.068)
1925-29	-0.615 (0.100)	0.316 (0.150)	-0.180 (0.085)
1930-34	-0.668 (0.201)	-0.439 (0.139)	-0.295 (0.133)
1935-39	-0.339 (0.210)	-0.367 (0.122)	-0.132 (0.082)
1940-44	-0.216 (0.128)	0.174 (0.106)	-0.010 (0.056)
1945-49	-0.187 (0.124)	0.311 (0.111)	0.019 (0.056)
1950-54	-0.364 (0.080)	0.311 (0.113)	-0.063 (0.057)
1955-59	-0.426 (0.074)	0.262 (0.115)	-0.098 (0.064)
1960-64	-0.438 (0.075)	0.229 (0.114)	-0.108 (0.066)
1965-69	-0.405 (0.075)	0.314 (0.117)	-0.082 (0.060)
1970-74	-0.231 (0.115)	0.245 (0.107)	-0.009 (0.055)
1975-80	-0.074 (0.175)	0.171 (0.122)	0.056 (0.063)
1920-80	-0.362 (0.094)	0.142 (0.103)	-0.082 (0.063)

no association between P_t and $\ell n\ E_t$. The specification of equation (9b) in Table 7 is the same as Lewis' (1964). In the estimates based on equation (9), the estimates of the parameters c_0, c_1, and c_2 change a good deal from equation to equation and often they are not significantly different from zero. The implied point estimates of $\bar{C}_t$ in Table 8 tend to move in the opposite direction from those for $\bar{B}_t$--that is, the manhours effects tend to be most negative in the early 1930's and, if negative, are close to zero in the late 1970's--but they are almost as frequently insignificantly different from zero as different from zero on two-tailed t-tests. Though there are slightly more negative values of $\bar{C}_t$ than positive values in columns (iv) and (v) of Table 8 and though the negative estimates are greater in absolute value than the positive estimates, they hardly constitute powerful evidence that unionism has depressed the level of manhours worked in sector a relative to sector b.

IV. Structural Equations

The estimates presented in the previous section have been labelled as those from "descriptive equations" in the sense that they have not made explicit use of structural labor supply and demand functions. In fact, Lewis did outline a structural model although he did not estimate all its parameters. The attractive feature of this structural approach is that, instead of the wage effect of unionism being measured independently of the effect of unionism on total manhours worked, the two effects are estimated as part of a single system of equations. It is important to observe that Lewis' structural supply and demand model

<u>assumes</u> that, if trade unions engineer a wage increase, unionized employers reduce manhours worked by moving back along their labor demand functions; wages and employment do not lie on the two parties' contract curve.

The relative demand function for labor is assumed to take the following simple form:

$$\ell n\ E_t = \delta_0 + \delta_1\ \ell n\ Q_t - \delta_2\ \ell n\ W_t$$

where both δ_1 and δ_2 are expected to be positive. The manhours supplied is given by the following equation:

$$\ell n\ E_t = \gamma_0 + \gamma_1\ \ell n\ W_t - (\delta_2 + \gamma_1)\bar{B}_t P_t \quad ,$$

where γ_1 is expected to be positive. In other words, at $\ell n\ W_t$, the number of manhours supplied in the absence of unionism would be $\gamma_0 + \gamma_1\ \ell n\ W_t$. However, in the presence of unionism, manhours actually used fall short of $\gamma_0 + \gamma_1\ \ell n\ W_t$ by an amount depending upon the size of the union sector, the magnitude of the relative wage effect, and the sum of the slopes of the log-linear supply and demand functions. In accordance with the findings of the previous section, specify $\bar{B}_t$ as a function of the inflation variable, $\ell n\ X_t$, and the unemployment rate, Z_t:

$$\bar{B}_t = \beta_0 + \beta_1\ \ell n\ X_t + \beta_2 Z_t \quad .$$

The equations determining the logarithms of equilibrium wage rates and equilibrium manhours, $\ell n\ W_t^e$ and $\ell n\ E_t^e$, respectively are as follows:

$$(10) \qquad \ell n\ W_t^e = \left(\frac{\delta_0 - \gamma_0}{\gamma_1 + \delta_2}\right) + \left(\frac{\delta_1}{\gamma_1 + \delta_2}\right) \ell n\ Q_t + \beta_0 P_t + \beta_1 (P_t\ \ell n\ X_t) + \beta_2 (P_t Z_t)$$

$$(11) \qquad \ell n\ E_t^e = \left(\frac{\gamma_0 \delta_2 + \gamma_1 \delta_0}{\gamma_1 + \delta_2}\right)$$

$$+ \left(\frac{\gamma_1 \delta_1}{\gamma_1 + \delta_2}\right) \ell n\ Q_t - \delta_2 \beta_0 P_t - \delta_2 \beta_1 (P_t\ \ell n\ X_t) - \delta_2 \beta_2 (P_t Z_t) \quad .$$

Consider now the relationship between these equilibrium relative wages and manhours on the one hand and the observed values of relative wages and manhours on the other hand. Lewis argued that, even in the absence of unionism, wage rates in each sector might not adjust to the same supply and demand conditions (that define their equilibrium values) at the same rate. In this event, transitory changes in the general price level ($\ell n\ X_t$) and in unemployment (Z_t) will induce movements in the ratio of wage rates in the two sectors relative to their equilibrium values.[25] If W_t denotes the ratio of wages observed in the two sectors during year t, then this argument concerning the relative movements in equilibirum and in actual wage rates may be expressed as follows:

$$(12) \qquad \ell n\ W_t = \alpha_1\ \ell n\ X_t + \alpha_2 Z_t + \ell n\ W_t^e + \varepsilon_{1t} \quad ,$$

where ε_{1t} is a normally distributed random variable. As for manhours, Lewis posited a simple lagged adjustment in the ratio of observed manhours in period t, E_t, to the ratio of equilibrium manhours in period t, E_t^e:[26]

$$\ln E_t = \lambda \ln E_t^e + (1 - \lambda)\ln E_{t-1} + \varepsilon_{2t} \quad , \tag{13}$$

where ε_{2t} is another normally distributed random variable. The error terms, ε_{1t} and ε_{2t}, are allowed to be contemporaneously correlated and each follows a simple, first-order, autoregressive process:

$$\varepsilon_{1t} = \rho_1 \varepsilon_{1t-1} + u_{1t} \tag{14}$$

$$\varepsilon_{2t} = \rho_2 \varepsilon_{2t-1} + u_{2t} \quad , \tag{15}$$

where u_{1t} and u_{2t} are joint, normally distributed, serially uncorrelated, random variables. If equations (12), (13), (14), and (15) are substituted into equations (10) and (11), we derive the following estimating equations for relative wages and relative employment:

$$\begin{aligned} \ln W_t = {} & \left(\frac{\delta_0 - \gamma_0}{\gamma_1 + \delta_2}\right) + \left(\frac{\delta_1}{\gamma_1 + \delta_2}\right)\ln Q_t + \alpha_1 \ln X_t + \alpha_2 Z_t + \beta_0 P_t \\ & + \beta_1 (P_t \ln X_t) + \beta_2 (P_t Z_t) + \rho_1 \varepsilon_{1t-1} + u_{1t} \end{aligned} \tag{16}$$

$$\begin{aligned} \ln E_t = {} & \lambda\left(\frac{\gamma_0 \delta_2 + \gamma_1 \delta_0}{\gamma_1 + \delta_2}\right) + \left(\frac{\lambda\gamma_1 \delta_1}{\gamma_1 + \delta_2}\right)\ln Q_t - \lambda\delta_2 \beta_0 P_t - \lambda\delta_2\beta_1(P_t \ln X_t) \\ & - \lambda\delta_2 \beta_2 (P_t Z_t) + (1 - \lambda)\ln E_{t-1} + \rho_2 \varepsilon_{2t-1} + u_{2t} \quad . \end{aligned} \tag{17}$$

The effect of unions on relative wages is given by a transformation of $\bar{B}_t = \beta_0 + \beta_1 \ln X_t + \beta_2 Z_t$ while the immediate impact of an increase in P_t on relative manhours is given by $\bar{C}_t = -\lambda\delta_2\bar{B}_t$. In this case, of course, $\bar{B}_t$ and $\bar{C}_t$ are estimated jointly and not independently of one another. The "descriptive equation" estimates of equations (16) and

(17) are given by equation (7b) of Table 5 and equation (9b) of Table 7 respectively. The maximum likelihood estimates of the parameters of equation (16) and (17) are as follows (with estimated standard errors in parentheses):[27]

$\hat{\delta}_0 = -0.479$ (0.943)	$\hat{\delta}_1 = 0.760$ (0.076)	$\hat{\delta}_2 = 0.705$ (0.287)	$\hat{\gamma}_0 = -0.983$ (1.011)
$\hat{\gamma}_1 = 3.213$ (0.799)	$\hat{\beta}_0 = 0.258$ (0.056)	$\hat{\beta}_1 = -2.590$ (0.514)	$\hat{\beta}_2 = -0.436$ (0.240)
$\hat{\alpha}_1 = 1.290$ (0.213)	$\hat{\alpha}_2 = 1.485$ (0.403)	$\hat{\lambda} = 1.025$ (0.047)	
$\hat{\rho}_1 = 0.523$ (0.110)	$\hat{\rho}_2 = 0.981$ (0.039)	log of likelihood function $= 284.6$.	

According to these estimates, neither the adjustment coefficient in manhours, λ, nor the serial correlation parameter in the manhours equation, ρ_2 , is significantly different from unity. The elasticity of the supply of relative manhours with respect to relative wages, γ_1, is estimated to be 3.21 while the elasticity of the demand for relative manhours with respect to relative wages, $-\delta_2$, is estimated to be -0.71.[28] As we found in estimating equation (7), the relative wage effect of unionism falls with increases in the unemployment rate (in other words, β_2 is estimated to have a negative, not a positive, impact upon $\bar{B}_t$) although $\bar{B}_t$ is estimated to be much more sensitive to movements in the inflation variable, $\ell n\ X_t$, than in unemployment.[29] The estimated values of $\bar{B}_t$ in particular subperiods and averaged over the entire 1920-80 period are given in column (iv) of Table 6. These estimated values are similar to, although a little lower than, those in the previous columns. Over the entire 1920-80 period, the relative wage effect of unionism (r) averaged 12.4 percent. The point estimates of $\bar{B}_t$ range widely over the sixty years although for many of these periods we are not in a position to reject the null hypothesis that $\bar{B}_t$ is zero. Because of our result that $\bar{B}_t$ moves inversely with the inflation variable, $\ln X_t$, once again our estimates imply lower values of $\bar{B}_t$ in the 1970's than in the 1950's and 1960's.

As for the relative manhours effects of unions, the estimates of $\bar{C}_t = -\lambda\delta_2\bar{B}_t$ are given in column (iii) of Table 8. By construction, these estimates of $\bar{C}_t$ are highest (in absolute value) in those subperiods when the relative wage effects are greatest.[30] However, what is most noticeable about these values of $\bar{C}_t$ is that only infrequently (1925-34) can we reject the null hypothesis that $\bar{C}_t$ is significantly different from zero. This result is consistent with the general inferences from estimating the "descriptive equations" (8) and (9): while negative impacts of unionism on relative manhours are often estimated, these effects are imprecisely estimated and confident inferences about $\bar{C}_t$ are unwarranted.

Thus, the inferences from these structural estimates are in line with those from the "descriptive equations" in Section III. There is strong evidence of a positive relative wage effect of unionism for much of the period from 1920 to 1980 and our best estimate is that it averaged around 12 percent. This estimate, however, is not measured with much confidence. There exists less compelling evidence of a negative effect of unionism on relative manhours worked and the null hypothesis that there exists no effect of unionism on relative manhours is frequently compatible with the evidence. The proposition that unionism depressed manhours worked in sector a below those worked in sector b can be accepted with some assurance only in the late 1920's and early 1930's.

V. Conclusions

The empirical analysis in this paper of the U.S. labor market from 1920 to 1980 leads us to conclude that, for most of this period, there is strong evidence of a positive impact of unionism on the average wage of union workers relative to the average wage of nonunion workers. The magnitude of this impact appears not to have been the same over the sixty years, but seems to have been larger than average in the early 1930's and smaller than average in the late 1940's and late 1970's. However, these relative wage effects are not measured with precision and a large number of different possible values for the effect of unions on relative wages are consistent with the evidence. These conclusions concerning relative wages (with the obvious exception concerning the effects in the 1970's) including the imprecision with which the relative wage impact of unionism is measured duplicate Lewis' findings.

The aggregate data imply a declining relative wage effect of unionism from the late 1960's to the late 1970's whereas the evidence from microeconomic studies tends not to accord with this finding. This may indicate that, owing perhaps to the growth of cost-of-living escalator clauses in collective bargaining contracts, the nature of the relationship associating the relative wage impact with macroeconomic variables (such as abnormal rates of inflation) changed in the 1960's and 1970's from the preceding decades. However, when we allowed for this possibility, our estimates yielded virtually the same inferences. Alternatively, this may indicate a mis-specification in the relationship for the entire period: if the sort of microeconomic studies that have been undertaken in the past ten years or so could be conducted with data on individuals from the 1930's and 1940's, perhaps these also would have produced estimates at variance with those from the aggregate data.[31/] It would be tempting to declare the results from the microeconomic studies as being the correct ones and to dismiss our estimates as being plagued with aggregation bias, but it should be noted that, if this is the case, then this means that the point estimates for the 1970's from the aggregate data tend towards underestimating the true impact of unionism on relative wages. In fact, in this literature on unions and wages, it is more frequently argued that the use of aggregate data produces an upward biased estimate of the true impact. We conclude that we do not know why the implications from fitting the equations to the aggregate time-series data do not agree with those from fitting wage equations to individual cross-section data in different years.

Although we are not at all confident about the precise magnitude of the effect of unionism on relative wages over the years from 1920 to 1980, we are satisfied that a positive effect existed for most of these years. By contrast, when it comes to the effect of unionism on relative manhours worked, we are not at all satisfied that the analysis of these data unambiguously points to a negative effect. We more frequently estimate negative effects of unionism on relative manhours worked than positive effects, but the estimates are disturbingly sensitive to small changes in the specification of the estimating equation. Even when negative manhour effects are calculated, our estimated standard errors do not permit us to make any useful confident statements and quite often we cannot reject the null hypothesis that unionism has not had any effect on relative manhours worked.

A fertile mind can think up a number of explanations for these results with respect to manhours worked, but one possibility is that Lewis' characterization of union wage behavior is not appropriate in all labor markets. According to Lewis' characterization, the unionized employer remains on his wage-taking labor demand function so the union's wage increase is achieved at the expense of a reduction in employment or manhours worked. This is a popular model of the way in which manhours are determined in unionized labor markets, but this popularity has not come about because it has been corroborated in a large number of empirical studies. On the contrary, there has been little attempt to test the empirical relevance of this model. In fact, by comparison with the number of studies designed to measure the relative wage impact of unionism, there have been few attempts to quantify the effects of unionism on manhours worked. This is surely an issue that deserves to be the subject of much more research.

In conclusion, we should like to emphasize that none of our results in this paper is at variance with Lewis' results and none of Lewis' judgments has required important qualification or modification. His research has been shown to possess a durability that few pieces of scholarship published in economics today are likely to have. Thus Reder's conjecture that, as a comprehensive analysis of the relative wage and employment effects of unionism, Lewis' work would be the best for a long time to come has been confirmed.

Footnotes

*/ We are indebted to John Raisian of the U.S. Bureau of Labor Statistics for providing us with the information to construct our series on the number of workers represented by trade unions. Support from the Sloan Foundation to the Department of Economics at Stanford University is gratefully acknowledged.

1/ To provide one comparison, consider the literature emanating from Harberger's (1962, 1966) estimates of the incidence of the corporation income tax which were published around the same time as Lewis' estimates of the economy-wide relative wage effects of unionism. Harberger's procedures and results immediately spawned a large literature: the model underlying his estimates was analyzed extensively, his numerical calculations were checked, and other economists determined whether his particular results could be produced by other methods. In fact, it transpired that there were both arithmetical and conceptual errors in Harberger's analysis (see Shoven, 1976). A discussion of this literature may be found in Atkinson and Stiglitz (1980).

2/ This is true, for instance, of the research investigating the consequences for measuring union relative wage effects of treating unionism as an endogenous variable. See Ashenfelter and Johnson (1972).

3/ A complete description of our data and of our methods in assembling the data are contained in the Appendix.

4/ We linked our series to Lewis' series P_2 (i.e., his adjusted estimates of the fraction unionized where the adjustment coefficient is 0.5).

5/ This statement rests on a comparison of the membership data and representation data collected by the CPSs for 1977 to 1980. These figures were divided by the number of full-time equivalent employees. The resulting difference between sector a and sector b in the fraction of workers who are union members is 0.255 in 1977, 0.243 in 1978, 0.244 in 1979, and 0.230 in 1980. By contrast, the difference between sector a and sector b in the fraction of workers represented by unions is 0.253 in 1977, 0.243 in 1978, 0.245 in 1979, and 0.230 in 1980. (The fractions presented in this footnote do not incorporate an adjustment to exclude an estimate of membership in employee associations. This is why neither of the series presented in this footnote corresponds to the entries for these years in column (vi) of Table 1. This is explained more fully in the Appendix.)

6/ This negative association after 1958 between the fraction unionized in sector a and the fraction unionized in sector b forms a marked contrast with the 1920-58 period: the simple correlation coefficient between P_a and P_b over the years 1920-58 is +0.972 while that during the period 1959-80 was -0.859. For the entire 1920-80 period the correlation coefficient between P_a and P_b is +0.743.

7/ The variables whose values for 1958 were changed from those used by Lewis were national income originating in the two sectors, total manhours worked per year, thousands of workers covered by collective bargaining agreements, full-time hours worked per worker per year, full-time equivalent employees, and the unemployment rate. The values of these variables that we used are given in the Appendix.

8/ If economic agents form their expectations of prices in an adaptive manner, then p_t^* measures the expected price level.

9/ When we embarked on our analysis, we employed two series for Z_t in the post-1958 period: the unemployment rate in the entire labor force and the unemployment rate of white men aged 35-44 years. The idea here was that, with a growing number of "secondary" workers in the labor force in the 1960's and 1970's, the overall unemployment rate might incorporate labor force composition effects that would render its movements less comparable with the movements in the unemployment rate before 1958. In fact, the results from using the unemployment rate of white men aged 35-44 years differed by only a very small amount from those using the overall unemployment rate so we retained the latter.

10/ The simple correlation coefficient between $\ell n\ X_t$ and Z_t is -0.690 for the years from 1920 to 1958, +0.554 from 1959 to 1980, and -0.644 for the entire period from 1920 to 1980.

11/ Equation (1) is simple identity: $\ell n\ W_{it} = P_{it}\ \ell n\ W_{it}^u$

$$+ (1 - P_{it})\ell n\ W_{it}^n = \ell n\ W_{it}^n + (\ell n\ W_{it}^u - \ell n\ W_{it}^n)P_{it}.$$

12/ This is, of course, only one possible meaning of "the relative wage effect of unionism." r does not measure the proportional difference between the wages of unionized workers and the wages these workers would have received in the absence of unionism. For a thorough treatment of different concepts, see Lewis (1963), Chapter II.

13/ For instance, the estimates of r in sector a range from 52.8 percent in equation (3b) to 14.5 percent in equation (3e) whereas the estimates of r in sector b range from 69.2 percent in equation (3c) to a negative value in equation (3e).

14/ The estimates in equations (6a) and (6b) are similar to Lewis' in his Table 58. We estimated to the entire 1920-80 period the specifications of the equations given in regression numbers 2 and 4 of Lewis' Table 60 and yielded almost identical results. Our one difference occurs when comparing the estimates of $\bar{B}$ in equations (6c) and (6d) of Table 4 with regression number 1 in Lewis' Table 62: our estimate of $\bar{B}$ for the 1920-80 period is almost double Lewis' for the 1920-58 period. In this instance, Lewis' unionism variable is not measured by his P_2 series that we have used. If we fit the specification given by equation (6e) to the 1920-58 period only, the coefficient on P_t is 0.160 with an estimated standard error of 0.055. This is much closer to Lewis' estimate.

15/ For equations (7a) and (7c), for instance, the calculated F statistics are 13.66 and 10.55 respectively compared with a critical value of about 3.2

16 For the record, the estimated standard errors of $\bar{B}$ corresponding to Lewis' Table 64 are as follows: 0.054 in 1920-24, 0.071 in 1925-29, 0.122 in 1930-34, 0.130 in 1935-39, 0.086 in 1940-44, 0.077 in 1945-49, 0.051 in 1950-54, and 0.050 in 1955-58.

17/ For instance, for the specifications given by equations (7a) and (7c) the calculated F statistics are 6.12 and 7.85 respectively while the critical F is about 3.2.

18/ In estimating this more general specification, we find that the partial effect of $\ell n\ X_t$ on $\bar{B}_t$ is less negative in the post-1958 period and the partial effect of Z_t on $\bar{B}_t$ is more negative in the post-1958 period by comparison with the pre-1958 years. With both $\ell n\ X_t$ and Z_t higher in the 1970's than in the preceding two decades, these two effects on $\bar{B}_t$ partially offset one another.

19/ The instruments for $\ell n\ Q_t$, P_t, $P_t\ \ell n\ X_t$, and $P_t Z_t$ were as follows: a constant term, $\ell n\ X_t$, Z_t, a quadratic time trend, the peak of the unemployment rate registered in the preceding recession, the fraction of Democrats in the House of Representatives, and a dummy variable taking the value of zero before 1932 and of unity from 1932 onwards. This last variable is designed to reflect the change in the attitude of government and the courts towards unionism as signalled first by the Norris-

La Guardia Act. The peak of the unemployment rate in the preceding recession and the fraction of Democrats in the House of Representatives were variables used (among others) by Ashenfelter and Pencavel (1969) to account for the level of unionism in the economy in year t conditional upon the level in the year $t - 1$.

20/ The average value of $\bar{B}_t$ for the entire 1920-80 period according to the specification in equation (7b) is 0.121 with an estimated standard error of 0.106.

21/ For instance, interacting this 1930's dummy variable with the unemployment rate according to the specification in equation (7a) and then testing for the joint significance of the coefficients on these interactions yields a calculated F statistic of 2.73 compared with a critical value of 3.18.

22/ We also considered the consequences of specifying a stochastic component in the determination of $\bar{B}_t$. If this were the case, then the error term in the estimating equation will not be homoskedastic, but will be a function of P_t^2. The results from addressing this issue--both those that assume $\bar{B}_t$ to be constant and those that relate $\bar{B}_t$ to $\ln X_t$ and Z_t--were similar to those reported in Tables 4 and 5.

23/ Our procedure here is different from Lewis' (1964). He took his estimated values of $\bar{B}_t$ from fitting the specification given by equation (7a) in Table 5 and treated them as if they were observations on a known, independent, variable. He regressed $\ln E_t$ on

$\ln Q_t$, $\ln E_{t-1}$, and the product of $\bar{B}_t$ and P_t and drew inferences about the effects of unionism on employment from the estimated coefficient on $\bar{B}_t P_t$.

24/ Lewis estimated his equation in first differences and in none of the specifications in Table 7 can the hypothesis that the first-order serial correlation parameter is unity be rejected.

25/ This is a drastic précis of Lewis' argument and we encourage the reader to consult pages 213-20 of his book.

26/ In fact, Lewis specified a different adjustment coefficient in the distributed lag on manhours demanded from that on manhours supplied. However, if the number of manhours supplied equals the number of manhours demanded, as he assumed, there is no way of identifying the two parameters separately.

27/ The Durbin-Watson statistic calculated from the values of $\hat{u}_{1t}$ is 1.70 while Durbin's h statistic calculated from the values of $\hat{u}_{2t}$ is 0.93.

28/ The estimates of this parameter δ_2 seem to be sensitive to the particular specification of the model. For instance, if no serial correlation in ε_{1t} and ε_{2t} is specified, the estimated value of δ_2 rises to 4.94 with an estimated standard error of 0.74 while λ falls to 0.37. Or, in estimating equations (16) and (17), if λ is constrained to unity, the estimated value of δ_2 becomes negative, -1.94, with an estimated standard error of 0.80.

29/ Evaluated at sample mean values of the variables, an increase in $\ell n\ X_t$ by one sample standard deviation (with Z_t constant at its mean) reduces the implied value of $\bar{B}_t$ from 0.117 to -0.050 while an increase in Z_t by one sample standard deviation (with $\ell n\ X_t$ constant at its mean) reduces the implied value of $\bar{B}_t$ from 0.117 to 0.092.

30/ The point estimates of $\bar{C}_t$ in column (iii) of Table 8 imply the following values for $\exp(\bar{C}_t)$ which is a natural index of the relative employment effects of unionism: 0.893 in 1920-24, 0.835 in 1925-29, 0.745 in 1930-34, 0.876 in 1935-39, 0.990 in 1940-44, 1.019 in 1945-49, 0.939 in 1950-54, 0.907 in 1955-59, 0.898 in 1960-64, 0.921 in 1965-69, 0.991 in 1970-74, 1.058 in 1975-80, and 0.921 for the entire 1920-80 period.

31/ Lewis' judgments about the relative wage impact of unionism over time were not drawn only from the aggregate data, but also from the examination of a number of studies of particular industries and occupations. These studies, however, had at their disposal fewer underlying observations than those that are available on magnetic tape and that have been the basis of so much work in the last decade.

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Appendix

Trade Union Representation

For the years from 1958 to 1980, the only continuous series on union membership is that from the Bureau of Labor Statistics' (BLS) survey of unions published in the *Directory of National Unions and Employee Associations*. Because this is a biennial survey, data at an industry level are available not each year, but every two years and values for intervening years must be estimated or interpolated. Even ignoring this problem, the *Directory* series has a number of features that render it suspect as reliable information on union membership by industrial sector.

First, although the aggregated data that are published distinguish between union membership within the United States and that abroad (primarily Canada), this distinction is not made at the industrial level. Moreover, Canadian membership has not remained a constant fraction of membership of all unions with headquarters in the United States: whereas in 1958 Canadian union membership represented 5.8 percent of total membership of unions with headquarters in the U.S., by 1978 this had risen to 7.1 percent.

A second problem with the membership data published in the *Directory* is that unions are thought to inflate their membership figures to present a slightly exaggerated impression of their size. Unemployed and retired members who are not paying their union dues are often kept on the union books. The retention of some of the unemployed on the membership rolls is especially relevant to a time-series on union membership given that the amplitude of business cycle movements in employment and working hours in highly unionized industries is much greater than in less unionized industries.

These two factors tend to make the _Directory_ figures an exaggerated estimate of the true union membership figures. A third factor pulls in the opposite direction and this concerns the distinction between union membership and the coverage of collective bargaining contracts negotiated by trade unions. The concept required in the analysis in the main body of the paper is the compensation of workers covered by union-negotiated contracts as a proportion of total employee compensation. Insofar as this is approximated more closely by the number of workers covered by collective bargaining contracts as a fraction of all employees, then it is coverage data rather than union membership data that are to be preferred. Now it is thought that, although the BLS surveys request information on union membership, some unions report their representation. Other unions do not, however, and in this event the reported membership figures may fall short of the coverage data. Some idea of the importance of this coverage-membership distinction can be gauged by the information in the first five columns of Table A which is taken from the May Current Population Surveys (CPS) for the years 1977, 1978, 1979, and 1980. In each of these years, the CPS asked the question "Are you a member of a union or employee association?" and, if the response was "no," then it proceeded with the question "Is your job covered or represented by a union?" The figures in the first five columns of Table A present the ratio of union

Table A

Union Coverage and Membership from the Current Population Surveys, 1977-80

	Ratio of coverage to membership					Thousands of workers represented by trade unions excluding an estimate of those represented by employee associations			
	1977	1978	1979	1980	Average	1977	1978	1979	1980
Mining	1.132	1.133	1.083	1.097	1.111	289.5	287.9	299.8	314.0
Construction	1.059	1.061	1.073	1.049	1.061	1,701.6	1,712.6	1,780.8	1,651.3
Manufacturing	1.059	1.074	1.081	1.079	1.073	7,583.2	7,629.6	8,181.2	7,308.6
Transport, Communications and Public Utilities	1.084	1.076	1.072	1.072	1.076	2,813.1	2,952.5	3,121.0	3,113.0
Wholesale Trade	1.089	1.102	1.142	1.109	1.110	444.5	428.4	471.4	431.8
Retail Trade	1.083	1.119	1.112	1.074	1.097	1,480.2	1,468.3	1,557.2	1,464.7
Finance, Insurance, and Real Estate	1.179	1.226	1.281	1.312	1.249	275.1	294.4	311.3	249.8
Services and Public Administration	1.217	1.211	1.199	1.204	1.208	4,637.6	4,749.7	5,203.9	5,319.8
Agriculture, Forestry and Fisheries	1.196	1.149	1.242	1.104	1.173	68.3	78.8	98.4	69.3
Sector a Industries	1.066	1.074	1.078	1.074	1.073	12,387.4	12,582.6	13,382.8	12,386.9
Sector b Industries	1.185	1.190	1.185	1.181	1.185	6,905.7	7,019.6	7,642.2	7,535.4
All Industries	1.114	1.120	1.122	1.119	1.119	19,293.1	19,602.2	21,025.0	19,922.3

representation to union membership by industry for the years 1977-80. The coverage-to-membership ratio is consistently higher in sector b than in sector a.

The information collected by the CPS represents the most accurate data on union coverage and so, in constructing our time series on the fraction unionized in sectors a and b over the years 1958-80, we started with these data for the years 1977-80 and worked backwards to the late 1950's. These CPS data in 1977-80 contain one problem for our purposes: in addition to membership in labor unions, they cover membership in employee associations, many of which do not engage in collective bargaining. These employee associations are in the public administration and service industries and fortunately the data collected by the BLS and published in the Directory distinguish between employee association membership and labor union membership. During the years 1977-80 only one Directory was published that provides this breakdown by industry between employee association membership and trade union membership and this is the Directory relating to 1978. According to the Directory, in 1978 there were 5,450 (thousands) members of labor unions and 2,634 (thousands) members of employee associations in the service industries and government combined. Or labor union membership represented 67.4 percent (that is, 5450/8084) of total membership in this broad industrial category. According to the 1978 Current Population Survey, total labor union and employee association membership in services and public administration was 5,818.2 (thousands) so that an estimate of labor union membership (excluding employee associations) in the 1978 CPS is 67.4 percent of 5,818.2 or 3,922.46 (thousands). As given in the second column of Table A, the ratio of union coverage to union membeship in 1978 in services and public administration was 1.2109 so that an estimate of the numbers of workers represented by trade unions (excluding employee associations) in this sector in 1978 is $1.2109 \times 3{,}922.46 = 4{,}749.7$ which is entered in the seventh column of Table A.

As has been mentioned, during the years 1977-80 only the Directory providing data for 1978 supplies information on employee association membership and labor union membership separately for different industries. Hence for the years 1977, 1979, and 1980 for services and public administration, we multiplied each of the membership figures reported in the Current Population Surveys (namely, 5,653.3 in 1977, 6,435.7 in 1979, and 6,555.0 in 1980) by 67.4 percent to derive an estimate of union membership (excluding employee association membership) and then inflated each figure by the ratio of representation to membership as given in the first, third, and fourth columns of Table A (namely, 1.217 in 1977, 1.199 in 1979, and 1.204 in 1980). This procedure yielded the figures in the sixth, eighth, and ninth columns of Table A, namely, 4,637.6 in 1977, 5.203.9 in 1979, and 5,319.8 in 1980. The numbers of workers represented by trade unions by broad industry groups as given by the May CPSs in the years from 1977 to 1980 are given in the final four columns of Table A.

Before 1977, the CPS asked just one question concerning union membership, namely, "Does . . . belong to a labor union?"--employee assocations are not mentioned nor is coverage. The membership data for the years 1976, 1975, 1974, 1973, 1970, and 1966 are given in Table B. The 1966 CPS provides information on the union membership of private workers only. The 1970 CPS gives data on the membership status of private workers and government workers separately and the ratio by industry of all union members to private union members was calculated from the 1970 data. This ratio was then applied to the private workers in 1966 in order to derive an estimate of all (private plus government) union members in 1966. Of the two columns under 1966 in Table B, the left column provides the data on private union memberships as given in the 1966 CPS while the right column represents the adjusted figures that incorporate an estimate of union membership among government workers.

In order to derive from these membership data in Table B an estimate of the numbers covered by union contracts, we multiplied the numbers in Table B by the ratios given in the fifth column of Table A (that is, 1.111 for mining, 1.061 for construction, and so on). These numbers in the fifth column of Table A represent the ratio by industry of union coverage to union membership averaged over the four years 1977-80. The resulting estimates by broad industry groups of workers covered by union contracts in 1966, 1970, and 1973-76 are given in Table C.

Table B

Thousands of Trade Union Members from the Current Population Surveys 1966, 1970, 1973-76

	1976	1975	1974	1973	1970	1966	
Mining	232.6	239.6	220.5	236.2	204	219	219
Construction	1457.4	1436.7	1652.8	1774.9	1948	1801	1911
Manufacturing	6990.9	6696.8	7820.1	7866.6	7600	8443	8506
Transport, Communications, and Public Utilities	2469.7	2455.0	2673.0	2536.4	2528	2183	2395
Wholesale Trade	386.6	375.6	399.7	388.1	345	1819	1829
Retail Trade	1338.5	1261.2	1320.1	1393.6	1363		
Finance, Insurance, and Real Estate	169.7	167.5	164.8	180.6	3155		
Services and Public Administration	4325.5	4101.4	3926.5	3643.4		978	1604
Agriculture, Forestry, and Fisheries	32.3	46.2	40.8	70.4	46		
Sector a Industries	11,150.6	10,828.1	12,366.4	12,414.1	12,280		13,031
Sector b Industries	6,252.6	5,951.9	5,851.9	5,676.1	4,909		3,433
All Industries	17,403.2	16,780.0	18,218.3	18,090.3	17,189		16,464

Table C

Estimates of the Thousands of Workers Represented by Trade Unions in 1966, 1970, and 1973-76

	1976	1975	1974	1973	1970	1966
Mining	258.4	266.2	245.0	262.4	226.6	243.3
Construction	1546.3	1524.3	1753.6	1883.2	2066.8	2027.6
Manufacturing	7501.2	7185.7	8391.0	8440.9	8154.8	9126.9
Transport, Communications, and Public Utilities	2657.4	2641.6	2876.1	2729.2	2720.1	2577.0
Wholesale Trade	429.1	416.9	443.7	430.8	383.0	2011.9
Retail Trade	1468.3	1383.5	1448.1	1528.8	1495.2	
Finance, Insurance, and Real Estate	212.0	209.2	205.8	225.6	3814.4	
Services and Public Administration	5225.2	4954.5	4743.2	4401.2		1939.2
Agriculture, Forestry, and Fisheries	37.9	54.2	47.9	82.6	54.0	
Sector a Industries	11,963.3	11,617.8	13,265.7	13,315.7	13,168.3	13,974.8
Sector b Industries	7,372.5	7,018.3	6,888.7	6,669.0	5,746.6	3,951.1
All Industries	19,335.8	18,636.1	20,154.4	19,984.7	18,914.9	17,925.9

All the information on union membership and coverage presented so far has been taken from the CPSs. However, these Surveys did not collect information on union membership or union coverage back to the late 1950's so, in order to extend the series back to those years, we must rely on the data collected every two years by the BLS and published in the Directory of National Unions and Employee Associations. For the years 1976, 1974, 1970, and 1966, a comparison of the CPS-based estimates of the number of workers covered by trade union contracts as given in Table C with trade union membership (excluding employee association membership) as published in the Directory is provided in Table D. Here the ratios of the representation figures in Table C to the membership figures published in the Directory show a systematic difference between sector a industries (mining, construction, manufacturing, transportation, communications, and public utilities) and sector b industries (wholesale and retail trade, finance, insurance, and real estate, services and public administration, and agriculture, forestry, and fisheries). Whereas the Directory figures on union membership tend to exceed (and, sometimes, exceed substantially) the numbers represented by unions in sector a, the membership figures in the Directory for sector b industries tend to understate the numbers represented by unions. Essentially, our procedure for generating a series on the numbers of workers represented by trade unions before 1966 in sectors a and b was to adjust the reported membership data as published in the

Table D

Ratio of Estimated Union Representation to Union Membership

	1976	1974	1970	1966
Mining	0.644	0.659	0.614	0.751
Construction	0.574	0.641	0.802	0.823
Manufacturing	0.876	0.918	0.889	1.041
Transport, Communications, and Public Utilities	0.827	0.883	0.828	0.766
Wholesale and Retail Trade	1.4441	1.424	1.213	1.487
Finance, Insurance, and Real Estate	4.327	6.431	1.027	
Services and Public Administration	1.079	1.014		0.680
Agriculture, Forestry, and Fisheries	0.790	1.331	2.250	

Directory by the estimates of the ratio of union representation to union membership as given in Table D. An account of our precise methods now follows.

Note that to this point, estimates of the numbers of workers represented by trade unions in sectors a and b have been derived for the years from 1973 to 1980 plus the years 1970 and 1966. Before turning to the pre-1966 data, consider filling the holes in the series between 1970 and 1973 and between 1966 and 1970. For the year 1972, we averaged the 1974 and 1970 ratios by industry of union representation to Directory union membership as given in Table D and then multiplied the resulting average for each industry by the published Directory membership in 1972. The number of workers represented by trade unions in 1972 as estimated this way is 12,994.9 (thousands) in sector a and 6,308.5 (thousands) in sector b. For 1971, each sector's representation figures for 1972 and 1970 were averaged: this yielded 13,081.6 (thousands) for sector a and 6,027.6 (thousands) for sector b. For the year 1968, each industry's ratio of representation to Directory membership for 1966 and 1970 (as given in Table D) was averaged and then this average ratio for each industry was then multiplied by the published Directory membership in 1968. The estimated number of workers represented by trade unions in 1968 is 13,827.3 (thousands) in sector a and 5,632.0 (thousands) in sector b. For 1969, each sector's representation figures for 1968 and 1970 were averaged, while for 1967 each sector's representation figures for 1968 were averaged: according to this procedure, sector a's union representation was 13,901.0 (thousands) in 1967 and 13,497.8 (thousands) in 1969 while the corresponding figures for sector b are 4,791.6 (thousands) in 1967 and 5,689.3 (thousands) in 1969.

Now consider constructing the series on the number of workers represented by trade unions in sectors a and b for the years before 1966. Our procedure was to adjust the membership data published every two years in the Directory by an estimate of the ratio of the number of workers represented by unions to the number of members reported in the Directory. For the manufacturing, transport, communications, and public utilities, and wholesale and retail trade industrial groups, the ratios we used to adjust the Directory membership figures were a simple average of the numbers in Table D for 1966, 1970, 1974, and 1976. These ratios were 0.931 for manufacturing, 0.826 for transport, communications, and public utilities, and 1.392 for wholesale and retail trade. The finance, insurance, and real estate membership data are not distinguished from the services and public administration data in the Directory for 1970 and 1966 so for these two industry groups the ratios we used to adjust the pre-1966 Directory data were computed as averages of the 1974 and 1976 numbers in Table D alone. This yields a ratio of representation to Directory membership of 5.379 for finance, insurance, and real estate and of 1.047 for services and public administration. Similarly, the union membership data for agriculture, forestry, and fisheries were not separately identified in the Directory membership for 1966 so for this industry the ratio of union representation to Directory membership was computed from the 1970, 1974, and 1976 entries of

Table D. This yields a ratio for this industrial group of 1.457. Finally, for mining and construction, we used the ratio of union representation to Directory membership as revealed by the 1966 Directory data alone, namely, 0.751 for mining and 0.823 for construction. The reason for not using the ratios averaged over the four years for these two industries in Table D (in which case the ratios would have been 0.667 for mining and 0.710 for construction) is that the numbers in the 1970's are heavily influenced by the growing numbers of retired United Mine Workers' members in the case of mining and by the growth in nonunion construction in the case of the construction industry. For the purpose of constructing a series on union representation in the early 1960's and late 1950's, we suspect that the 1966 data on the ratio of representation to Directory membership are likely to yield a more accurate estimate of union representation for these two industries than would an average of the 1966, 1970, 1974, and 1976 ratios.

With these ratios of union representation to Directory membership, we then went to the biennial issues of the Directory from 1958 to 1966 and multiplied these ratios by each industry's reported membership figure in each year. The results when aggregated across industries to form sectors a and b are shown in columns (i) and (ii) of Table E. Now because the ratios of union representation to Directory membership that we used to adjust the Directory membership data were not based on the 1966 information alone (except for mining and construction), there is no guarantee that the estimates of numbers represented by unions in 1966 from so adjusting the Directory membership will equal the estimates of

Table E

Estimated Thousands of Workers Represented by Trade Unions, 1956-66

	Sector a Column (i)	Sector b Column (ii)	Sector a Column (iii)	Sector b Column (iv)	Total Column (v)
1956	13,238.1	3,851.5	14,001.9	2,954.1	16,956.0
1958	12,953.9	4,195.3	13,701.3	3,217.8	16,919.1
1960	12,999.5	4,110.6	13,749.6	3,152.8	16,902.4
1962	12,486.4	4,130.8	13,206.9	3,168.3	16,375.2
1964	12,538.5	4,645.1	13,262.0	3,562.8	16,824.8
1966	13,212.9	5,153.4	13,975.3	3,952.7	17,928.0

NOTES: Each entry in Column (iii) is 1.0577 times the corresponding entry in Column (i) and each entry in Column (iv) is 0.7670 times the corresponding entry in Column (ii). Column (v) is the sum of Columns (iii) and (iv) numbers represented by unions in 1966 from the CPSs. Indeed, as given in Table C, the latter are 13,974.8 (thousands) for sector a and 3,951.1 (thousands) for sector b, whereas the implied figures for union representation in 1966 by adjusting the Directory memberships data (as given in columns (i) and (ii) of Table E) are 13,212.9 (thousands) for sector a and 5,153.4 (thousands) for sector b. Consequently, we then returned to the pre-1966 data and readjusted sector a's series by multiplying each year's observation by 1.0577 (= 13,974.8/13,212.9) and readjusted sector b's series by multiplying each year's observation by 0.7670 (= 3,951.1/5,153.4). The resulting series is given for sector a in column (iii) of Table E and for sector b in column (iv) of Table E. Finally, for the intervening odd numbered years between 1956 and 1966, the series on numbers of workers represented by trade unions in columns (iii) and (iv) of Table E were linearly interpolated. The numbers so constructed for the years from 1958 to 1980 are given in Table 1: the estimated numbers of workers covered by union contracts in sector a are in column (i), those in sector b are in column (ii), and those in the entire economy are in column (iii); column (iv) gives the numbers in column (i) divided by the number of full-time equivalent employees in sector a; column (v) gives the numbers in column (ii) divided by the number of full-time equivalent employees in sector b; column (vi) is the difference between the numbers in columns (iv) and those in column (v); and the numbers in column (vii) are those in column (iii) divided by the total number of full-time equivalent employees.

How accurate are these estimates of the numbers represented by trade unions and how do the estimates in this paper compare with other estimates? Those estimates computed from the CPSs in the 1970's are likely to be quite accurate. Those inferred from the published Directory membership data are likely to contain nontrivial errors of measurement. We know of no-one else who has tried to calculate a time-series of the number of workers covered by collective bargaining contracts in the late 1950's and 1960's. Leo Troy (1969) has carefully compiled labor union membership data on a consistent basis for the years 1961-66 and the total numbers of workers covered by union contracts as calculated in this paper average 2.8 percent higher than his membership data. A differential such as this is to be expected insofar as union coverage estimates tend to exceed union membership estimates. If we were to extend our procedures back to 1956, our implied estimates of the fraction of workers represented by trade unions are those given in columns (i) and (ii) of Table F. Lewis' estimates of the extent of union membership are given in columns (iv) and (v), (vii) and (viii), and (x) and (xi). Again, for both sectors a and b, Lewis' membership estimates never exceed and are usually smaller than our representation estimates. The variable used in most of the regression analysis, however, is the arithmetic difference between the fraction unionized in sectors a and b and a comparison of our estimates of this difference in column (iii) with Lewis' estimates in columns (vi), (ix), and (xii) will reveal smaller discrepancies between our estimates and Lewis'. It would be absurd to pretend that our unionism data so constructed do not

Table F

A Comparison of Our Fraction Represented by Unions with Lewis' Union Membership Proportions, 1956-58

	Estimated by Procedures Used in this Paper			Lewis' Unadjusted Estimates			Lewis' Adjusted Estimates with Adjustment Coefficient of 0.7			Lewis' Adjusted Estimates with Adjustment Coefficient of 0.5		
	Sector a Column (i)	Sector b Column (ii)	Difference Column (iii)	Sector a Column (iv)	Sector b Column (v)	Difference Column (vi)	Sector a Column (vii)	Sector b Column (viii)	Difference Column (ix)	Sector a Column (x)	Sector b Column (xi)	Difference Column (xii)
1956	0.569	0.106	0.463	0.545	0.094	0.451	0.548	0.096	0.452	0.552	0.098	0.454
1957	0.566	0.108	0.458	0.557	0.095	0.462	0.561	0.096	0.465	0.566	0.097	0.469
1958	0.606	0.113	0.494	0.576	0.095	0.481	0.565	0.096	0.469	0.551	0.097	0.454

NOTES: The estimates of the fraction of full-time equivalent employees represented by trade unions in Columns (i), (ii), and (iii) are those derived by the procedues outlined in this paper. The other estimates in this table are taken from Table 51, page 199, of Lewis (1963).

contain serious defects. They certainly contain errors of unknown magnitude although it is doubtful whether these errors are greater than those underlying Lewis' series. Lewis, of course, was fully aware of the shortcomings of his own data and, indeed, we may conclude by recommending the reader consult Lewis' own discussion of these issues--first pages 258-76 of his book and then pages 198-201.

Apart from the CPS tapes, the published documents consulted in preparing these series on union representation are as follows. The issues of the *Directory of National Unions and Employee Associations* consisted of Bureau of Labor Statistics Bulletins 2079, 2044, 1937, 1750, 1665, 1596, 1493, 1395, 1320, 1267, and 1222. The *Directory* for 1973 appears not to have been issued as a BLS Bulletin. Also consulted were U.S. Department of Commerce, *Population Characteristics: Labor Union Membership in 1966*, Current Population Reports, Series P-20, No. 216, March 1971; U.S. Department of Labor, Bureau of Labor Statistics, *Selected Earnings and Demographic Characteristics of Union Members, 1970*, Report 417, 1972; and Courtney D. Gifford, ed., *Directory of U.S. Labor Organizations 1982-83 Edition*, Bureau of National Affairs, Washington, D.C. 1982.

National Income

The national income originating from each of the ten major industry divisions is to be found in Table 6.3B of the Department of Commerce's National Income and Product Accounts. The figures for 1957-75 are from [A] and those for 1976-80 are from [B]. Sector a's figure for each year is derived by adding the national income originating in the following industry divisions: mining; construction; manufacturing; and transportation and public utilities. Sector b's figure for each year is derived by adding the national income originating in the following industry groups: wholesale trade; retail trade; finance, insurance, and real estate; services; government and government enterprises except the military; and agriculture, forestry, and fisheries. To exclude national income originating in the military, the total compensation of military employees (Table 6.5B) was subtracted from the government and government enterprises entry.

The figures thus constructed were linked to Lewis' series as they appear in his Table 59, page 211, although we linked our numbers to his 1957 figure rather than to his 1958 figure. That is, we replaced his 1958 indices of 4.286 for sector a and 3.923 for sector b with 4.297 and 3.969 respectively. The reason for this modification was that our procedures yielded estimates that replicated Lewis' series almost perfectly up to and including 1957, but they diverged slightly for 1958. The figures we use reflect the latest adjustments and corrections to the data and so we are inclined to use our figures for 1958 over Lewis'. The difference, however, is only slight. Our series for 1957-80 are given in columns (vii) and (viii) of Table G and they link up with Lewis' series in columns (1) and (2) of his Table 59, page 211.

Average Annual Full-Time Compensation

For each major industry division (that is, mining; construction; manufacturing; transportation and public utilities; wholesale trade; retail trade and automobile services; finance, insurance, and real

Table G

Observations on Other Variables Used in the Analysis, 1958-80

	Average Annual Full-Time Compensation (Dollars)		Full-Time Hours Worked per Worker per Year		Total Manhours Worked		Index Numbers of National Income		Ratio of Actual To Expected Prices	Unemployment Rate
	Sector a Column (i)	Sector b Column (ii)	Sector a Column (iii)	Sector b Column (iv)	Sector a Column (v)	Sector b Column (vi)	Sector a Column (vii)	Sector b Column (viii)	Column (ix)	Column (x)
1958	5,415.0	3,889.0	1,998.1	2,048.5	1.095	1.255	4.297	3.969	1.042	0.0633
1959	5,720.8	4,120.9	2,028.2	2,051.1	1.154	1.289	4.841	4.219	1.043	0.0512
1960	5,910.0	4,252.9	2,007.4	2,057.2	1.147	1.333	4.915	4.436	1.039	0.0512
1961	6,094.7	4,430.6	2,010.2	2,049.5	1.120	1.343	4.957	4.648	1.031	0.0623
1962	6,360.8	4,635.1	2,024.0	2,050.3	1.159	1.372	5.331	4.941	1.032	0.0512
1963	6,589.1	4,830.4	2,030.7	2,050.9	1.171	1.404	5.641	5.200	1.030	0.0530
1964	6,931.2	5,051.9	2,042.4	2,046.9	1.196	1.437	6.056	5.571	1.030	0.0484
1965	7,126.3	5,252.8	2,040.0	2,017.7	1.253	1.483	6.656	6.038	1.033	0.0419
1966	7,465.6	5,530.8	2,040.0	2,023.7	1.325	1.534	7.315	6.612	1.042	0.0353
1967	7,763.9	5,816.8	2,015.9	1,994.5	1.322	1.569	7.504	7.122	1.046	0.0353
1968	8,343.3	6,239.8	2,008.8	1,977.5	1.341	1.610	8.162	7.800	1.058	0.0335
1969	8,840.1	6,719.4	2,009.7	1,981.6	1.373	1.664	8.495	8.691	1.071	0.0326
1970	9,504.9	7,338.4	1,979.1	1,970.9	1.308	1.687	8.674	9.220	1.081	0.0456
1971	10,194.1	7,757.3	1,976.7	1,971.3	1.269	1.720	9.217	10.045	1.085	0.0549
1972	10,919.1	8,293.3	1,991.9	1,970.4	1.314	1.778	10.222	11.067	1.082	0.0521
1973	11,723.6	8,840.7	1,997.4	1,954.0	1.384	1.843	11.495	12.492	1.090	0.0456
1974	12,756.5	9,516.2	1,968.8	1,928.5	1.363	1.869	12.263	13.486	1.113	0.0521
1975	14,137.9	10,348.8	1,957.0	1,931.6	1.242	1.888	13.004	14.787	1.131	0.0791
1976	15,381.7	11,156.1	1,962.6	1,919.2	1.288	1.935	14.874	16.218	1.118	0.0716
1977	16,643.9	11,952.7	1,966.1	1,914.6	1.343	1.992	16.778	18.172	1.113	0.0651
1978	17,994.7	12,832.9	1,968.5	1,896.8	1.416	2.072	18.993	20.682	1.120	0.0558
1979	19,943.1	13,872.7	1,963.6	1,890.0	1.462	2.134	20.924	23.229	1.133	0.0540
1980	21,670.3	15,191.1	1,947.5	1,875.8	1.405	2.154	22.088	25.409	1.145	0.0660

estate; services; government and government enterprises except the military; agriculture, forestry, and fisheries), average annual full-time compensation was computed from national income data by adding to "wages and salaries per full-time equivalent employee" (Table 6.9B of the National Income statistical tables) an estimate of supplementary compensation per full-time equivalent employee. The latter, following Lewis, was calculated by subtracting all annual wages and salaries (Table 6.6B) from annual total compensation of employees (Table 6.5B) and then dividing by the number of full-time equivalent employees each year (Table 6.8B). (In each case, the figures for the military were subtracted from the entire government sector.) Figures for sectors a and b were calculated as fixed weighted averages of these industry estimates where the weights are the industry's annual full-time equivalent employment relative to the group's total full-time equivalent employment over the entire period 1958-80. (Note that these fixed weights are not the same as Lewis' fixed weights which were the average relative employment of each industry over the period 1929-57.) The series for the years 1958 to 1975 were taken from [A] below and the series for the years 1976 to 1980 were taken from [B] below. With sector a and b compensation thus constructed, these series were linked to Lewis' series in Table 54, page 204, of his book. The series from 1958 to 1980 is given in columns (i) and (ii) of Table G.

Full-Time Hours Worked per Worker per Year

For each of the ten major industry divisions, annual full-time hours worked per employee were calculated as the total number of hours worked per year by full-time and part-time employees (Table 6.10 of the National Income statistical tables) divided by the number of full-time equivalent employees per year (Table 6.8B). In each year the figures for the military were subtracted from the government's sector. The sector averages were fixed weighted averages of the industry estimates where the weights are the industry's annual full-time equivalent employment relative to the group's total full-time equivalent employment over the entire period 1958-80. The sources were [C] below for the years 1957-72, [A] (except for the government sector) for 1973-74, [D] for 1975-78, and [B] (except for the government sector) for 1979 where Table 6.12 provided the hours data. Hours worked and full-time equivalent employees for the military and for the entire government sector in the years 1973 and 1974 were taken from the issues of the Survey of Current Business for July 1977 and 1978 respectively. For the government sector in 1979, the hours worked of the military seems not to be available in the regularly published series so for this year it was inferred by assuming it was the same proportion of the hours worked in the entire government sector as it was in 1978.

The sources cited do not provide data for 1980. For this year we turned to the household data from the CPS which provides information on average weekly hours worked (of those at work). We assumed that for each industry the ratio of the Department of Commerce numbers to the CPS numbers in 1979 remained the same in 1980 so, with the latter data available in 1980, we can infer the corresponding Department of Commerce figures for hours worked. The CPS data are published in Employment and

Earnings, January 1980 and January 1981 (Tables 34 and 32 in both issues). Because data for mining are not supplied from the CPS, we used the establishment data (Employment and Earnings, January 1982, Table 3, page 198).

With the sector a and b hours worked per employee thus constructed, these series were linked with Lewis' series in Table 56, page 206, of his book. The link was made with Lewis' observations on average hours worked for 1957 not for 1958. That is, we replaced his 1958 indices of 1.088 for sector a and 1.245 for sector b with 1.095 and 1.255 respectively. (In fact, our revised values for 1958 yield a ratio of 0.873 which is virtually the same as Lewis' ratio of 0.874.) The reason for replacing Lewis' values with ours is that we have been able to make use of revisions in the data series that presumably reflect improvements in the underlying data. Our series on hours worked are listed in columns (iii) and (iv) of Table G.

Total Manhours Worked

For each industry division, total manhours worked is the product of (a) full-time hours per year per full-time equivalent employee and (b) the number of full-time equivalent employees. The construction of (a) has already been discussed. Data on full-time equivalent employees by industry is from Table 6.8B of [A] for 1957-76 and of [B] for 1977-80. After each industry's total manhours worked figure is formed, the sector figure is derived by summing over the constituent industry figures. The implied index numbers of total manhours worked per year are given in columns (v) and (vi) of Table G; these link up with those given in columns (4) and (5) of Lewis' Table 59, page 211, of his book.

Ratio of Actual to Expected Price Level

The index used for prices is the implicit price deflator for net national product available from Table 7.6 of issues of the National Income and Product Accounts. The series on "expected" prices is calculated in the same way as Lewis formed it, namely, as a weighted average of current and past prices with weights declining exponentially from the current date to earlier dates. The ratio of actual to expected price levels that we constructed is given in column (ix) of Table G and it links up with Lewis' series in column (2) of Table 61, page 218, of his book.

Unemployment Rate

The series on the unemployment rate was taken from the Employment and Training Report of the President, 1981, Table A-1, page 119, and was linked up in 1957 with Lewis' series in column (1) of Table 61, page 218 of his book. The series is presented in column (x) of Table G.

References for Appendix

The list of the National Product Accounts publications consulted is as follows:

[A] U.S. Department of Commerce, Bureau of Economic Analysis, The National Income and Product Accounts of the United States, 1929-76, Statistical Tables, A Supplement to the Survey of Current Business, September 1981

[B] U.S. Department of Commerce, Bureau of Economic Analysis, Survey of Current Business, Vol. 62, No. 7, July 1982.

[C] U.S. Department of Commerce, Bureau of Economic Analysis, The National Income and Product Accounts of the United States, 1929-74, Statistical Tables.

[D] U.S. Department of Commerce, Bureau of Economic Analysis, Survey of Current Business, Vol. 59, No. 7, July 1979.